Discussion Paper No. 00-04

Returns to Education in West Germany

An Empirical Assessment

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Economic Research

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Nontechnical summary

This paper analyses the developments in the returns to education in West Germany from 1984 to 1997. Simple estimates of the returns to schooling showed that the returns to one additional year of schooling in terms of wages have remained remarkably stable since the mid-1980s. Women have significantly higher returns to schooling (about 10%) than men (about 8%). We tested the robustness of these results in various ways, taking into account possible cohort effects, the choice of the sample, the definition of the human capital variables and different estimation methods.

Firstly, the analysis showed that the returns to education are not constant over the life-cycle, especially for women. Evaluating the returns to schooling for different cohorts at the same age shows that a significant decline in the returns to education across cohorts is observable at age 30 to 39, and this decline is particularly pronounced for women since 1994. At the middle of the career (age 40-49), however, we found evidence for slightly increasing (men) or constant (women) returns across birth cohorts. Finally, at an older age (50 to 60), the returns to education are lower for younger cohorts, particularly for women beginning in 1994.

Secondly, we examined whether differences in the specification of human capital variables in the wage equation could alter the estimates of the returns to education. Departing from the quantitative measure of education, we analysed the returns to educational degrees. The higher the degree, the higher the wage premium. However, when we correct for the different length of studies associated with the various degrees, the master degree yields the highest return. A downward trend is observable for the return to the master degree and to the high school diploma at the end of the period. Our tests for alternative specifications of the other variables showed that the level of the return to schooling is quite robust. Only when using age instead of labour market experience, the returns to education are somewhat lower. Female returns to education are somewhat sensitive to the inclusion of additional variables designed to capture previous non-employment or to control for the industrial and regional structure, but the difference from the simple specification is not really important.

Thirdly, the analysis reveals huge differences across subgroups of workers. The returns to education are much higher for part-time than for full-time working women (12-13% versus 8%). Taking full-time workers only, there are no significant gender differences: the return amounts to about 8% for both genders. Moreover, the differences between public and private sectors are also remarkable. In the public sector, female returns to schooling have increased somewhat (from about 9% to 10%) over the period, whereas male returns have slightly decreased (from about 8% to 7%). As a result, female returns are significantly higher than male returns. In the private sector, however, female returns decrease much stronger than male ones. Consequently, women have now lower returns to schooling than men.

Finally, we examined methodological issues and focussed on the impact of possible estimation biases on the level of the returns to education. The results do not point to the presence of significant selectivity and endogeneity biases.

On the whole, the simple estimates proved quite robust towards specification and estimation method. However, the overall assessment hides some more complex developments, i.e. huge differences between subgroups of workers or conjunction of time, life-cycle and cohort effects. Thus, studies on the returns to education in West Germany should be interpreted very carefully and one should be aware of the implications of the specific framework adopted.

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Abstract: This paper analyses the developments in the returns to education in West Germany for the period from 1984 to 1997. Based on simple Mincer-type wage equations, we estimate a return of about 8% for men and 10% for women, and these returns have remained remarkably stable over the period. On the basis of more differentiated specifications of wage equations, we find evidence for the presence of cohort effects, in addition to time and life-cycle effects. Furthermore, the analysis shows that the choice of the sample of observation plays a crucial role. Indeed, huge differences exist between part-timers and full-timers, as well as between private and public sectors. Full-time working women have similar returns to schooling than men, and if female returns are declining and have become lower than male returns in the private sector, they are rather increasing and are higher than male ones in the public sector. Moreover, not all education degrees yield the same annual return. If one accounts for the different lengths of studies, the master craftsman degree yields the highest return. However, the estimates proved rather robust towards the specification of the wage equation and the estimation method.

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1 Introduction

Following human capital theory (Becker 1964, Mincer 1974), engaging in further education can be seen as an investment which first entails costs – at least foregone earnings –, then yields a return in the form of increased future wages. So far, no consensus view has emerged from the empirical literature on returns to education regarding the level of these returns and their developments over the past decades in Germany (see Lauer and Steiner 1999 for a review of the empirical literature on this topic). As a matter of fact, most recent studies for Germany do not focus explicitly on the relationship between education and wages, but rather on the wage structure or the wage distribution with respect to industrial sector (e.g. Fitzenberger and Kurz 1996, Dustman and Van Soest 1998), gender (e.g. Gerlach 1987, Prey 1999) or region (Giles et al. 1998), especially in comparisons of Western and Eastern Germany (e.g. Steiner and Wagner 1998, Franz and Steiner 1999). In these studies, education, one of many other human capital variables, is treated as a control variable rather than as of interest *per se*. Only a few studies explicitly analyse the distribution of earnings in connection with the qualification structure (e.g. Bellman, Reinberg and Tessaring 1994, Weißhuhn and Clement 1983 for the 1970s and the beginning of the 1980s). Moreover, the various studies differ with respect to the sample and the period observed, the specification and the estimation methods. As a result, they are hardly comparable and the results only hold within the specific framework adopted.

The aim of this paper is twofold. First, it aims at providing an overall assessment of the developments in the returns to education in West Germany over the period from 1984 to 1997. We restrict our analysis to West Germany, because changes in the East German wage distribution after unification differ fundamentally from West German wage developments in the 1990s (see Steiner and Wagner 1998, Franz and Steiner 1999). Second, it aims at examining the extent to which the overall assessment depends on the specific framework adopted. To this end, we test for the sensitivity of the results in alternative contexts. By doing this, we hope to identify essential elements to be kept in mind for the interpretation of any study on the returns to education.

The paper is organised as follows. Section 2 explores the basic relationship between education and wages. Following a brief descriptive overview, the estimation of standard Mincer wage equations provides the basis for bringing out the trend in the returns to education in West Germany since the mid-eighties. Section 3 explores the sensitivity of these basic results to different specifications of the wage equation. First of all, the possible presence of cohort effects in the returns to education is analysed (section 3.1). Indeed, developments of the returns to education may result from genuine time effects, but also from cohort effects or life-cycle effects. Section 3.2 deals with the impact of the specification of the estimated wage function, focussing on the definition of the education variable - in quantitative terms (years of schooling) or qualitative terms (education levels) - and on the definition of the other human capital variables. Furthermore, the sensitivity of the returns to the choice of the sample observed will be examined (section 3.3). In particular, it may be of interest to distinguish between full-time and part-time workers, between private and public sectors. Finally, the impact of the choice of the estimation method on the level of the returns to education will be analysed, particularly in selectivity issues and possible endogeneity of the schooling variable (section 3.4). Similar estimations were also conducted in fourteen other countries within the framework of a project funded by the European Commission[1]. For comparability reasons, in all partner countries, the extensions in the sensitivity analysis are separate variants of the reference Mincer model and do not build upon each other.

[1] TSER project „Public Funding and the Returns to Education“.

2 Basic relationship between education and wages

2.1 Descriptive overview

Table 1 shows the qualification structure of West-German employees[2] aged between 30 and 60 in 1984 and in 1997[3]. The proportion of employees with no degree at all is close to zero. The great majority (about two thirds) of West-German employees holds a low or intermediate school degree (*Hauptschul-* or *Realschulabschluß*) assorted with an apprenticeship or a master craftsman degree or equivalent (*Meister*, *Fachschule*, *Beamten-ausbildung*, *Gesundheitsausbildung*). Conversely, very few have a high school degree alone (*Abitur* or *Fachhochschulreife*). The bulk of high school leavers pursue their studies by completing an apprenticeship (particularly women) or higher education (particularly men). At the tertiary level, two thirds of all employees have completed university or equivalent institutions, and only one third the more practically oriented and short-track higher technical colleges (*Fachhochschulen*). The proportion of tertiary level graduates has increased from about 12% to some 18% between 1984 and 1997. Similarly, the proportion of high school graduates, especially those holding an additional vocational degree, as well as the share of master craftsmen has increased significantly, whereas lower qualified employees have become comparatively fewer. This points out a shift of the qualification structure upwards over the period. Women are overrepresented in the lowest educational categories (no vocational degree) and underrepresented at the tertiary level. However, the proportion of women with no vocational degree has decreased strongly.

Table 1: Qualification structure 1984 und 1997 (%)

	1984		**1997**		**Total**	
	Men	Women	Men	Women	1984	1997
No degree	0.0	0.3	0.0	0.1	0.1	0.1
Low or interm. school	10.0	29.6	6.0	14.5	16.9	9.4
Apprenticeship	55.6	51.9	47.4	55.6	54.3	50.7
Master	16.1	7.4	16.9	9.4	13.0	13.9
High school	0.1	0.3	1.1	0.8	0.1	1.0
High school+ appr./master	4.2	2.0	6.6	6.8	3.5	6.7
Higher tech. college	5.4	2.6	8.2	3.1	4.4	6.1
University	8.6	6.1	13.4	9.0	7.7	11.6

Source: GSOEP 1984-97, own calculations.

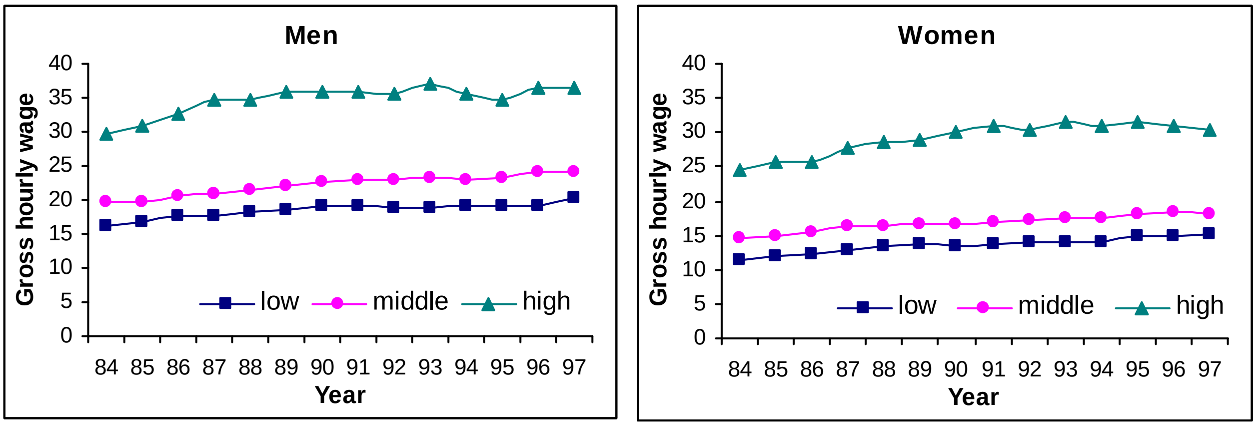
In the following figures, West-German employees aged between 30 and 60 have been grouped into three broad educational categories:

- Low education level (“low“): no degree or only a low or intermediate school degree.
- Intermediate education level (“middle”): apprenticehip, master, high school with or without apprenticeship/master.
- High education level (“high”): higher technical college or university.

[2] This study is based on data from the 14 waves of the German Socio-Economic Panel (GSOEP) currently available (see in Lauer and Steiner 1999 a description of the GSOEP). For the purpose of the analysis, the sample was restricted to West-German citizens. The self-employed, pensioners, military personnel, people engaged in education or training were also excluded.

[3] See the description of the German education system in Annex 1.

Figure 1: Average real gross hourly wage by education level



Source: GSOEP 1984-97, own calculations.

The real gross hourly wage is obtained by dividing the monthly nominal gross wage in the month preceding the date of interview by the number of hours worked, and deflating it by the consumer price index. As can be seen from Figure 1, there is a huge gap between gross hourly wages of tertiary level graduates and the rest. Real gross hourly wages of unskilled and skilled workers have developed in a quite parallel way over the period. Table 2 shows that the wage increase was somewhat stronger for the lower qualified (25% for men and even about 33% for women). The wage increase was also stronger for women, especially at lower qualification levels. Men earn more than women at all education levels, but the difference is narrower at higher education levels: whereas low-educated men earn about 35% more than their female counterparts, male tertiary level graduates earn "only" 20% more than female graduates. The gender wage differential seems to have shrunk, especially at lower qualification levels.

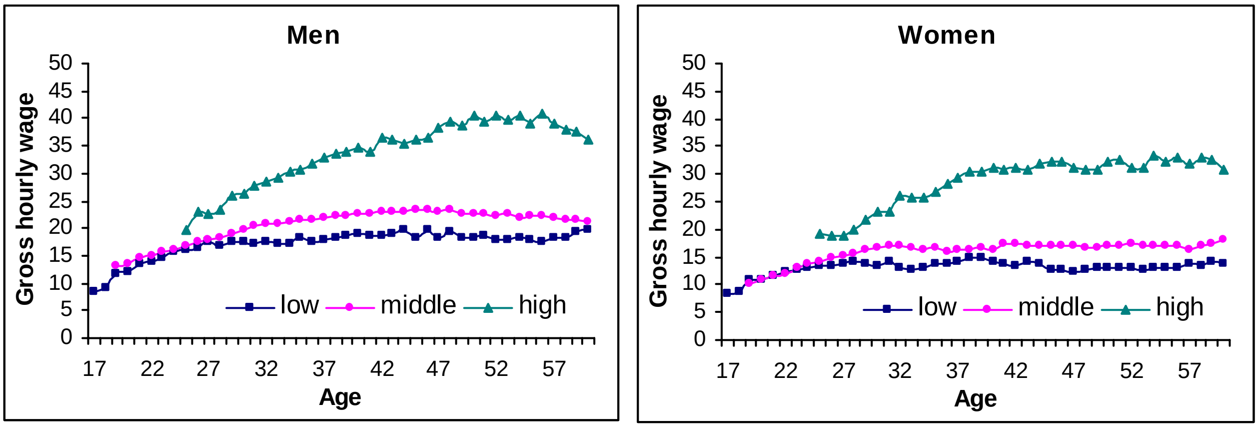
Table 2: Increase in real gross hourly wage and gender wage gap[1)] by education level 1984 and 1997 (%)

		Low	Middle	High
Wage increase (%)	Men	25.3	22.3	23.1
	Women	32.6	24.2	23.7
Gender wage gap (%)	1984	41.5	34.6	21.7
	1997	33.8	32.5	20.1

Source: GSOEP 1984-97, own calculations.
Note: 1) Difference between male and female wages as a percentage of female wages.

These stylised facts tend to corroborate the view that education, especially higher education, yields a positive return. However, educational background is not the only factor influencing labour productivity and, hence, wages. Non school-based human capital is also expected to raise wages. In particular, labour market experience is likely to be a strong determinant of wages. Figure 2 shows the pattern of wages over the life-cycle, differentiated by education levels. Given that the length of studies differs across the education levels, the working careers associated with the various education levels start at different ages. That is the reason why the sample was reduced to individuals aged between 30 and 60. Not doing so would result in overrepresenting the lower educated. Even at the beginning of their career, tertiary level graduates earn more than lower educated, although at this stage, the lower educated have a couple of years of working experience behind them. At all education levels, a sort of concave shape can be observed: in a first stage, wages increase faster, then more slowly[4]. The turning point is located at different ages across genders and across education levels. Generally speaking, it is earlier for women. This is probably due to the fact that women traditionally interrupt their working career owing to family duties and thus accumulate less human capital. Another observation is that the turning point is later for higher education levels. This implies that the educational wage premium becomes larger along with age.

Figure 2: Age-wage profiles by qualification level (DM)



Source: GSOEP 1984-97, own calculations.

Having described the most important stylised facts concerning the relationship between wages and education in Germany, we now turn to a more detailed analysis of the returns to education in West Germany on the basis of empirical wage equations. Such an analysis makes it possible to assess the returns to education while controlling for differing labour market experience.

2.2 Simple estimates of the returns to education

The standard approach to the estimation of the returns to education dates back to Becker (1964) and Mincer (1974) and has its roots in the neoclassical theory. The basic assumption is that an individual's earnings reflect its labour productivity and that investment in human capital in the form of foregone earnings in the past pays off in the form of higher wages in the future (see Card (1999) for an extensive discussion of the human capital theory and its empirical implementation). Starting from this, Mincer developed a theoretical model from which he derived the following wage equation:

$$ln(Wage_i) = \alpha_0 + \alpha_1\, Schooling_i + \alpha_2\, Experience_i + \alpha_3\, Experience_i^2 + u_i$$

with $ln(Wage)$ = log of gross hourly wage
$Schooling$ = years of education
$Experience$ = years of labour market experience
u = error term
i = index for individuals
α_j = coefficients to be estimated, j=0,1,2,3.

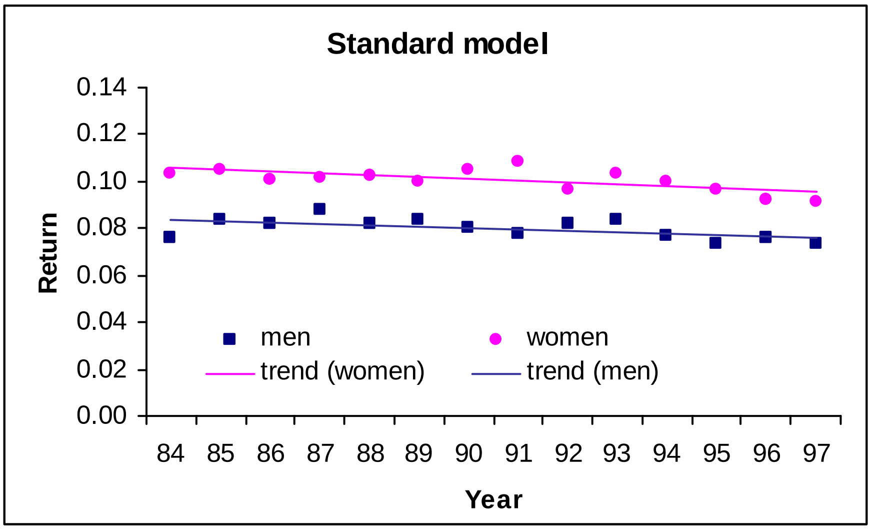
[4] In the empirical literature, age-earnings profiles are most often aggregated across education levels, typically from the age of 20. This leads to a stronger, but somewhat misleading, concavity: the strong increase at the beginning of the life-cycle is mainly due to the fact that, at young ages, the more educated have not yet arrived on the labour market. When they start working (typically between age 25 and 30), the average wage level increases sharply.

In principle, this equation can easily be estimated empirically. The semi-loglinear specification of the wage equation relates to some functional form assumptions underlying the theoretical derivation of Mincer's wage function. More importantly, it also corresponds to the observed distribution of wages. While the schooling variable proxies human capital acquired through formal education, labour market experience is a proxy for human capital acquired on-the-job. The inclusion of labour market experience in linear and quadratic form also relates to the particular derivation of Mincer's wage function and is designed to capture the concave shape of wage-experience profiles.

The error term captures all factors other than schooling and labour market experience affecting individual wages. Typically, the error term is assumed to be normally distributed and uncorrelated with the human capital variables as well as between individuals and, in case the wage equation is estimated on panel data, across time. Given this assumption is true and assuming the wage equation is correctly specified, OLS yields unbiased parameter estimates. Under the assumption that there are no other costs associated with education than foregone earnings, the estimated coefficient of the schooling variable (α_1) directly measures the returns to one additional year of education in terms of (log) wages, and the coefficients associated with the experience variables (α_2 and α_3) measure the return to labour market experience. To allow for changes in coefficients over time, the wage equation is estimated separately for each wave of the GSOEP. Since we are particularly interested in gender differences in estimated returns to human capital, we estimate the wage equation for men and women separately.

The dependent variable is real gross hourly wage. Years of schooling are derived from information on the highest completed educational or vocational degree by attaching an average number of years to standardised educational levels. Following usual practice, labour market experience is defined as potential experience, that is, age minus years of schooling minus school starting age (6 years). To avoid an overrepresentation of the lower educated in the sample due to earlier career starting age, we focus on the population aged 30 to 60. All estimated coefficients are significant at the 1% level, and this simple model explains about 35-40% of the wage variance in the male sample and about 30% in the female sample.

Figure 3: Return[1)] to education 1984-97



Source: GSOEP 1984-97, own calculations.
Note: 1) Coefficient of the schooling variable α_1. To obtain the return in percent, compute [exp(α_1)-1] ×100.

The estimated returns to schooling are quite stable over time (the slight downward trend observable in Figure 3 is hardly significant). Hence, we have pooled all waves of the GSOEP and re-estimated the wage equations on the pooled sample, which yields estimates of the average returns to schooling and experience over the observation period.[5].

Table 3: Estimation results of the wage equation on the pooled sample - Dependent variable: log gross hourly wage

	Men		**Women**		**t-test[2)]**
	Coefficient[1)]	Std. err.	Coefficient	Std. err.	Men=Women
Schooling	0.080	0.001	0.100	0.002	-11.18
Experience	0.041	0.001	0.025	0.002	6.14
Experience²/100	-0.068	0.002	-0.043	0.004	-5.27
constant	1.495	0.022	1.163	0.041	7.15
Observations	16,988		10,240		
R^2	0.39		0.34		

Source: GSOEP 1984-97, own calculations.

Notes: 1) To obtain the returns in percent, compute [exp(α_i)-1] ×100. 2) H_0: $\text{Return}_{men} = \text{Return}_{women}$. t-statistics $= (b_2 - b_1) / \sqrt{(se_1)^2 + (se_2)^2}$. If $|t|>1.96$ (resp. 2.58, 1.65), then the hypothesis of equality of the coefficients is rejected at a significance level of 5% (resp. 1%, 10%).

We find a coefficient of 0.08 for men and of 0.10 for women, which corresponds to a return of about 8.3% (=exp[(0.08)-1] ×100) for men and 10.5% for women per additional year of schooling. The t-tests show that the gender differences with respect to the returns to schooling are highly significant. The returns to labour market experience cannot be directly read off the table. From the specification of the wage equation above, the effect of labour market experience on the log wage is given by $\hat{\alpha}_2 + 2\hat{\alpha}_3 Experience$, where $\hat{\alpha}_2$ and $\hat{\alpha}_3$ are the estimated parameters of the linear and quadratic experience terms in the wage equation. Therefore, the return to one additional year of labour market experience depends on the level of experience.

Table 4: Return[1)] to labour market experience at different experience levels[2)]

Years of labour market experience	**Men**		**Women**	
	Return	Std. error	Return	Std. error
5	0.034	0.0006	0.021	0.0008
10	0.027	0.0004	0.017	0.0006
15	0.021	0.0003	0.012	0.0004
20	0.014	0.0002	0.008	0.0003
25	0.007	0.0002	0.004	0.0003
30	0.000	0.0003	-0.001	0.0004

Source: GSOEP 1984-97, own calculations.

Note: 1) $\hat{\alpha}_2 + 2\hat{\alpha}_3 Experience$. To obtain the return in percent, compute [exp($\hat{\alpha}_2 + 2\hat{\alpha}_3 Experience$)-1] ×100.
2) The computation of the standard errors is based on the "delta" method: For a given level of X , let the function $h(\alpha_2, \alpha_3) = \alpha_2 + 2\alpha_3 X$. The variance of the estimate $\hat{\alpha}$ obtained from $h(\hat{\alpha}_2, \hat{\alpha}_3)$ is $Var(\hat{\alpha}) = AVA'$, with the matrix $A = [\partial h / \partial \alpha_2, \partial h / \partial \alpha_3]$ and V = variance-covariance matrix of the estimates $\hat{a}_2$ and $\hat{a}_3$.

[5] This is fully legitimate for women, since further F-tests applied to interaction terms between the explanatory variables and year dummies in a regression on the pooled sample showed that the differences in the coefficients over time are statistically insignificant. For men, however, the hypothesis of joint insignificance of the interaction terms was not clearly rejected (all interactions were jointly insignificant at the 5% level, but not the interactions with the schooling variable alone). Thus, the coefficients drawn from the pooled sample have to be interpreted as average values.

Table 4 shows that, first, the marginal return to experience decreases substantially as the level of labour market experience increases. This confirms the observations from the descriptive overview. Secondly, at all levels of experience, the marginal return to experience is higher for men than for women. For example, the return to an additional year of labour market experience after 5 years of experience amounts to about 3.5% for men and 2.1% for women. After 15 years of experience, the marginal return to experience is approximately 2% for men and 1.7% for women, after 30 years it approaches zero.

3 Sensitivity analysis

In the previous section, we followed the traditional approach for estimating empirical wage equations. This quite parsimonious model, applied to a specific sample, produced two main results regarding the returns to education: they have remained remarkably stable over time, and they are higher for women (about 10%) than for men (about 8%). Now, we explore the robustness of these results by estimating and testing alternative specifications of the wage function. We concentrate on four main questions:

- Are there cohort effects in the returns to education (section 3.1)?
- Does the level of return depend on the specification of human capital variables (section 3.2)?
- Are there notable differences across groups of workers in the returns to education? In particular, do the observed trend and gender differences depend on the sample chosen (section 3.3)?
- Is such a standard model subject to some estimation bias which might influence the level of return to education obtained (section 3.4)?

3.1 Cohort and life-cycle effects

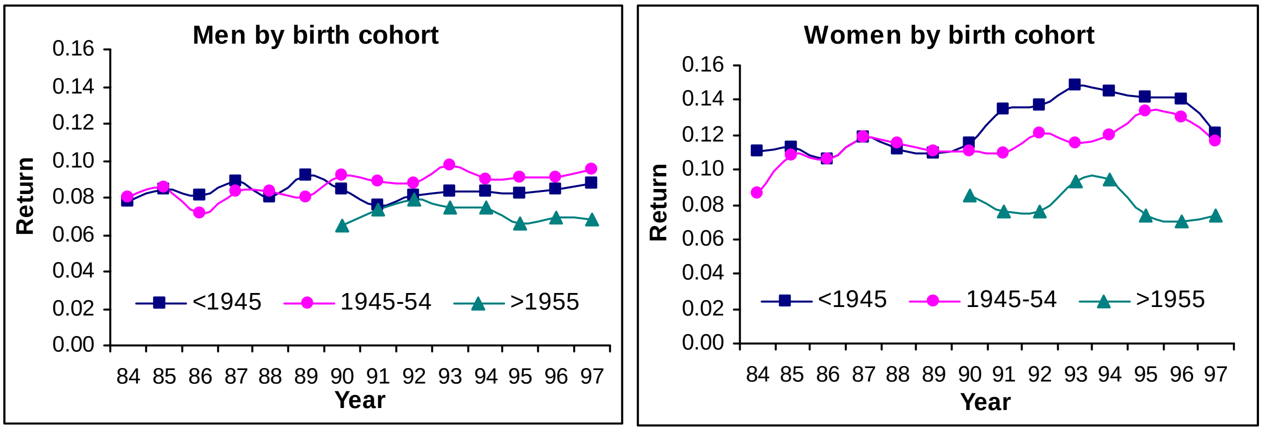
In the reference model, we estimated the developments over time in average returns to education across all cohorts. This section aims at distinguishing between different effects which might influence the developments in the return to education:

- *Time effects*. The economic environment has changed during the 1980s and the 1990s, with phases of growth and recession, rising unemployment etc. By estimating the wage equations year by year, the standard approach intends to identify time effects.
- *Cohort effects*. It may also be of importance in which year the observed individuals were born: e.g. are they part of the baby-boomers generation? Are there some noticeable differences in the returns to education across birth cohorts, in particular, do younger birth cohorts have lower or higher returns to education?
- *Life-cycle effects*. The age of the observed individuals is essential: at which stage of the life-cycle are the observed individuals? Do the return to education vary over the life-cycle?

These effects are difficult to disentangle (see Heckman and Robb 1985). In fact, it is impossible to isolate them perfectly one from another, because the date of birth, the year and the age are inextricably linked in a linear relationship. Thus, it is impossible to observe two different birth cohorts at the same age and at the same year. However, Figure 4 and Figure 5, by adopting different approaches, provide useful hints concerning possible cohort and life-cycle effects. Ideally, we would like to be able to examine each cohort over the whole life-cycle. In that case, we could run the regression separately for each cohort, while controlling for life-cycle effects by including a term for labour market experience and its square. The problem is that if we can control for time developments within a cohort through year dummies, for instance, we cannot control for the fact that the cohorts would be observed in different periods in time. Here, the problem is a little bit different: since we only have 14 years of observation, we can only catch one particular phase of the life-cycle for each cohort we want to observe. If the returns to education turn out not to be constant over the life-cycle, this may bias the estimated returns. Therefore, it seems useful to also analyse differences across cohorts at the same age.

In Figure 4, we split the sample according to the birth cohort. In order to have a sufficient number of observations, we distinguish between three cohorts: people born before 1945, people born between 1945 and 1955 and those born in 1955 or after. By doing so, we can compare the returns to education across different cohorts at a given year (reading the graph vertically) and follow the developments in the returns to education over a specific phase of their respective life-cycle which corresponds to their age during the period 1984 to 1997 (reading the graph horizontally).

Figure 4: Return[1)] to education 1984-97 by birth cohort



Source: GSOEP 1984-97, own calculations.

Note: 1) Coefficient of the schooling variable α_1. To obtain the return in percent, compute $[\exp(\alpha_1)\text{-}1] \times 100$.

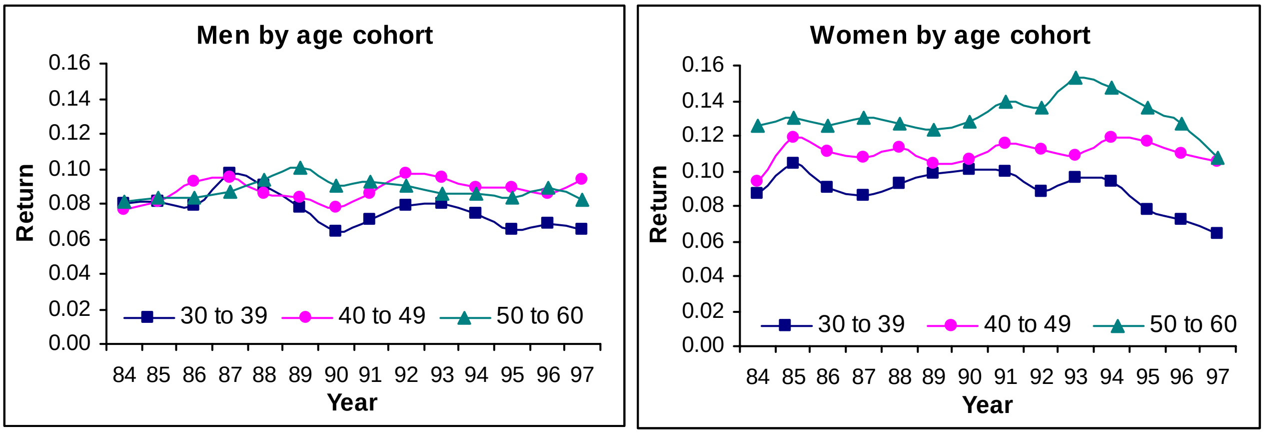
For men, the returns to education of the older cohort (born before 1945) and of the middle one (born between 1945 and 1954) have increased over the period, respectively from 8% (=[exp(0.077)-1]×100) to 9.2% and from 8.3% to 10%. For women, the older cohort enjoy significantly higher returns to education than the middle one after 1990. The increase in the returns to education of the middle cohort is very pronounced, from 9% to 12%. Since we focus on the population aged between 30 and 60, the younger cohort is only represented in the second half of the observation period. For both men and women, the youngest cohort, aged 35 to 42 during this period, has the lowest returns to schooling, and this is particularly pronounced for women. Interestingly, the returns of the younger cohort are not increasing over the life cycle.

There are several interpretations for these developments. One interpretation would be that they depict the developments in the returns over the life-cycle: at the beginning of the life-cycle, male returns to education stagnate, in the middle of the life-cycle, they increase at a faster rate than at the end of their career. For women, the returns first stagnate, then increase strongly until the end of the career. The specific pattern for women could be attributable to the fact that many women interrupt their career in their prime years for family reasons and have not yet reaped the full benefit of their education at the first stage of their career, but catch up later. Following this life-cycle interpretation, men or women of the youngest cohort are assumed to experience the same developments in their returns when they attain a later stage of their working career as the older cohorts experience now. However, this interpretation is only valid if there are no cohort effects.

Another key interpretation – which does not exclude the first one, but complements it - would be that the return of the youngest cohort is lower not because it is at an earlier stage of the life-cycle, but because the returns have decreased across cohorts. This would mean that the returns of the youngest cohort will not be as high as that of the middle cohort when they reach the same stage of the life-cycle. An argument for this interpretation is that there are some jumps over the cohorts which are not explainable by life-cycle developments. For instance, the return to schooling of the youngest female cohort in 1997 should correspond more or less to the return of the 10-year older cohort 10 years earlier, i.e. in 1987. However the latter is 12,5% and thus much higher than the return of the former group of 7.7%. For men, the same is observable, although to a lesser extent. Therefore, we suppose that there are, in addition to life-cycle effects, cohort effects affecting the returns to education downwards, especially for women.

In order to account for cohorts being at different stages of their life-cycle and to complement the analysis, we compare in Figure 5 the returns to education for different cohorts at the same age. We selected three age groups: from 30 to 39, from 40 to 49 and from 50 to 60 years. Hence, the age cohort 30 to 39 in 1984 was born between 1945 and 1954, the age cohort 30-39 in 1985 corresponds to the birth cohort 1946-1955, and so on. Similarly, the age cohort 40-49 in 1984 is the birth cohort 1935-44, and the age cohort 40-49 in 1985 is the birth cohort 1936-45. Reading the graph horizontally, we can compare the returns to education for older (left) and younger (right) birth cohorts at a given age. The defined birth cohorts are overlapping, so that the values reported are moving averages. Reading the graph vertically, you get the differences in the average return across age cohorts at a given year, i.e. across different birth cohorts.

Figure 5: Return[1)] to education 1984-97 by age cohorts



Source: GSOEP 1984-97, own calculations.

Note: 1) Coefficient of the schooling variable α_1. To obtain the return in percent, compute [exp(α_1)-1] $\times$100.

For men, the birth cohort aged 50 to 60 old in 1997 obtains about the same level of return to education as the cohort aged 50 to 60 in 1984 (about 8%). This general statement hides a first upward trend across birth cohorts with a peak at about 10.6% ([exp(0.100)-1]$\times$100) in 1989 (birth cohort 1929-1939), followed by a downward trend for the younger cohorts between 50 and 60 years old. At age 40-49, however, some increase in favour of younger cohorts is observable: the return to education increased from less than 8% for the age cohort 40-49 in 1984 (birth cohort 1935-1944) to about 9.8% for the age cohort 40-49 in 1997 (birth cohort 1948-1957). At an earlier stage of the career (30-39 year-olds), the returns are clearly lower for younger cohorts: the return to education decreased from about 10.6% for the cohort aged 30 to 39 in 1987 (birth cohort 1948-57) to 6.7% for the cohort aged 30 to 39 in 1997 (birth cohort 1958-67).

For women, the differences between the age groups are much more pronounced: at each given year, it is apparent that the older the age group, the higher the return. This means that the increase in the returns to education over the life cycle is much more marked for women than for men. This confirms the observations drawn from Figure 4. For women, the returns to education at age 40-49 remain fairly constant between the cohort aged 40-49 in 1984 and the cohort aged 40-49 fourteen years later. However, at an older age, the returns first remain stable, then increase significantly, and from 1993 the returns to education dropped sharply from 16.6% down to around 11.3%. At an earlier stage of the working career, a large decline across cohorts is observable, from about 11% to 6.6%. This decline is particularly pronounced from 1994 onwards, i.e. for birth cohorts born after 1955-64. The figure also shows that in all age groups, the returns to education have declined since 1994, even if the decline is less pronounced for the middle-age group. This suggests that, in addition to cohort effects, time effects may have played a role in the decline of the female return to education.

3.2 Choice of the specification

This section examines whether alternative specifications of the standard human capital wage function modify the findings. We focus on two issues: the definition of the education variable itself, and the specification of the other variables of the wage equation.

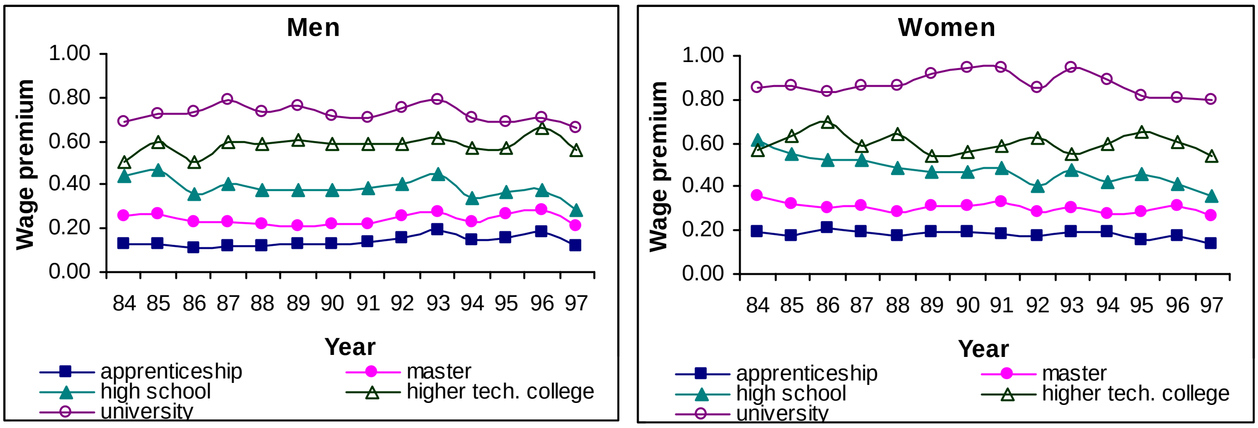
3.2.1 Education levels instead of years of schooling

The reference model measures education in a quantitative way, through the number of years of schooling. This approach implicitly suggests that one additional year of schooling, whatever the current level of education is, yields the same return. This may not be the case if, for instance, completed degrees rather than years of schooling as such are valued in the labour market. In this section, we depart from the quantitative specification and allow education to affect wages in a non-linear way by including dummy variables for the highest completed educational or vocational degree in the earnings function. Here, we use the same categories as in the descriptive overview in Part 2.1. The reference group consists of individuals without any degree or with only a low or intermediate school degree (*Hauptschul-* or *Realschul-abschluß*). Holders of high school degree, with or without an additional vocational degree, have been grouped together. Hence, the wage equation we estimate here for each year from 1984 to 1997 is the following:

$$ln(Wage_i) = \alpha_0 + \alpha_1\, Apprenticeship_i + \alpha_2\, Master_i + \alpha_3\, High\ school_i + \alpha_4\, Higher\ technical\ college_i + \alpha_5\, University_i + \alpha_6\, Experience_i + \alpha_7\, Experience_i^2 + u_i,$$

where the educational variables are dummies taking on the value 1 if the individual has the corresponding education level and 0 otherwise. Figure 6 reports the coefficients estimated from the regression. They represent the wage premium associated with the different education degrees compared to the reference group (no degree or only a low/intermediate school degree). However, in order to obtain an idea of the effective yearly return corresponding to each type of educational degree, one should take into account that the completion of the different educational degrees requires different durations of studies. The higher educated start working later and this reduces their effective return to education, since they incur a longer period of foregone earnings. The coefficients reported in Table 5 have been corrected for differing lengths of studies[6] and thus represent the yearly returns to the degrees considered.

Figure 6: Wage premia[1)] associated with education degrees 1984-97



Source: GSOEP 1984-97, own calculations.

Note: 1) Coefficients α_i of the i dummies for educational degrees. To obtain the return in percent, compute $[\exp(\alpha_i)-1] \times 100$.

Comparing Figure 6 and Table 5 allows us to disentangle the trend in the relative labour market value of diplomas in terms of wages from the trend in the private return to these diplomas, which is also influenced by developments in the relative length of studies[7]. Not surprisingly, the higher the educational level, the higher the wage premium. Particularly for women, there seems to be a high bonus in pursuing university studies. However, once controlled for the different length of studies, the hierarchy changes. The master degree yields by far the highest return, both for men and women. This is attributable to its short length of studies, compared, for instance, to tertiary level studies. However, a decreasing trend is observable in the returns to the master degree. For men, this is entirely due to the fact that the study duration gap between master craftsmen and the reference group has increased somewhat, whereas the wage premium do not decrease. For women, however, this is both the outcome of a declining labour market valuation of female masters and of a change in the duration gap. A striking feature is that the wage premium for employees with a high school diploma (with or without an additional vocational degree) has decreased sharply. This is particularly true for women, for whom the returns to high school dropped from 12.5% in 1984 to 7.4% in 1997. For men, this decrease only started in 1993, but is also quite strong (from 8-9% to less than 6%). This decrease is only the result of a declining labour market valuation of this degree, and not of changes in the duration gap. As a result, a high school degree yields the lowest return at the end of the period, for both men and women. Having this degree in addition to an apprenticeship does not seem to bring any further return. The returns to education are higher for women than for men in all educational categories except for holders of a higher technical degree. The trend in the return to higher technical college is constant over time for men, and slightly declining for women. This is mainly due to the fact that more and more women complete an apprenticeship prior to higher technical college studies, which increases the duration of studies. At the very end of the period, the returns to higher technical college are similar for both genders.

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[6] The information about actual schooling years is not available, therefore, it was constructed by adding up average number of years needed to complete certain educational levels. We prefer to present the results in table form rather than by a graph because of the superposing of the curves.

[7] This is true under the assumption that foregone earnings are the only costs associated with schooling.

Table 5: Return[1)] to educational degrees 1984-97

	Men					Women				
	Apprent.	Master	High school	High. tech. coll.	Univ.	Apprent.	Master	High school	High. tech. coll.	Univ.
1984	0.079	0.121	0.086	0.079	0.075	0.113	0.155	0.118	0.104	0.096
1985	0.080	0.125	0.092	0.093	0.080	0.104	0.142	0.109	0.114	0.096
1986	0.067	0.109	0.070	0.078	0.080	0.122	0.132	0.103	0.118	0.093
1987	0.075	0.109	0.083	0.092	0.086	0.113	0.140	0.104	0.105	0.096
1988	0.069	0.103	0.075	0.092	0.078	0.098	0.120	0.095	0.108	0.096
1989	0.075	0.098	0.077	0.090	0.082	0.113	0.130	0.093	0.091	0.103
1990	0.076	0.100	0.076	0.089	0.077	0.109	0.130	0.091	0.091	0.106
1991	0.075	0.098	0.077	0.088	0.076	0.106	0.139	0.097	0.093	0.105
1992	0.089	0.111	0.081	0.090	0.082	0.100	0.120	0.081	0.099	0.095
1993	0.107	0.122	0.090	0.092	0.085	0.109	0.124	0.095	0.085	0.105
1994	0.080	0.102	0.067	0.086	0.077	0.109	0.115	0.084	0.096	0.099
1995	0.084	0.113	0.072	0.083	0.073	0.088	0.121	0.089	0.102	0.091
1996	0.099	0.119	0.074	0.096	0.076	0.097	0.128	0.082	0.096	0.090
1997	0.068	0.091	0.057	0.083	0.072	0.080	0.114	0.072	0.083	0.088

Source: GSOEP 1984-97, own calculations.

Note: 1) $return_i = \alpha_i / (d_i - d_{ref.})$, with d_i = number of years required to complete degree i. To obtain the return in percent, compute $[\exp(r_i)-1] \times 100$.

3.2.2 Various measures of non school-based human capital

In the reference model, we used the typical potential experience, i.e. age minus years of schooling minus school starting age (6 years), as a measure for non school-based human capital. This measure has the advantage of controlling for the individual's age as well as for the fact that potential working experience is expected to differ depending on the length of education. However, potential labour market experience is only a rough indicator for actual experience. Typically, it is a poor indicator for female labour market experience, since women tend to interrupt their working career to devote their time to family duties. Moreover, labour market experience is only one measure of an individual's productivity. A series of other variables may also be relevant. For this reason, we tested the effects of the following specifications of the wage equation on the returns to education:

- Reference model: This is the standard model, with *potential experience* and its square.
- Model 1: Instead of potential experience, we use *age* and its square.
- Model 2: Instead of potential experience, we use *actual experience* and its square.
- Model 3: Instead of potential experience, we use *actual full-time experience* and its square.
- Model 4: In addition to potential experience, we include a variable for the duration of *previous non-employment* and its square.
- Model 5: In addition to model 4, we include a variable for *previous unemployment* and its square.
- Model 6: In addition to model 5, we include a variable for *tenure* with the current employer and its square, and a dummy variable for *part-time* work.
- Model 7: In addition to model 6, we include further control variables for *firm size, industry branch* and *region of residence*.

The variables designed to account for previous non-employment, previous unemployment and actual experience are constructed on the basis of the retrospective data on the employment status of the individuals from the age of 15 years contained in the GSOEP. Actual experience is constructed by adding the length of all full-time and part-time employment spells observed in the potential working career, i.e. after completion of initial training. Similarly, all spells of unemployment are added to build the "previous unemployment" variable. The variable for previous non-employment measures cumulated non-employment during the working career and is obtained by adding up all spells of unemployment, housekeeping, military service and other non-employment activities. The variables themselves, as well as a quadratic term and interaction terms, are included in the function in order to allow for non-linear experience effects. We include tenure with the current employer as a proxy for specific human capital acquired within the firm. Direct information about tenure is available in the GSOEP. Dummy variables are built for firm size, industry and region. For this part of the analysis, we concentrate on a single year, namely 1995, since we are mostly interested in the effect of adding further variables on the level of the returns to education rather than on the trend.

In Table 6, the results of the schooling coefficients as well as of t-tests are reported; the latter are included to check whether the alternative specifications differ with respect to the return to education from the reference specification where potential experience is used.

Table 6: Sensitivity of returns[1)] to education to different specifications of human capital variables other than schooling (1995)

	Men					Women				
	Return	Std. err.	Obs.	R^2	t-test[2)]	Return	Std. err.	Obs.	R^2	t-test[2)]
Ref. model	0.073	0.003	1138	0.33	Ref.	0.097	0.006	722	0.32	Ref.
Model 1	0.064	0.003	1138	0.30	-2.09	0.088	0.005	722	0.31	-1.09
Model 2	0.071	0.003	1138	0.34	-0.60	0.098	0.005	722	0.34	0.14
Model 3	0.072	0.003	1138	0.35	-0.36	0.094	0.005	722	0.35	-0.40
Model 4	0.071	0.003	1138	0.35	-0.50	0.091	0.006	722	0.37	-0.69
Model 5	0.070	0.003	1138	0.37	-0.85	0.096	0.006	722	0.37	-0.13
Model 6	0.068	0.003	1137	0.38	-1.26	0.091	0.006	720	0.38	-0.70
Model 7	0.067	0.003	1135	0.47	-1.40	0.079	0.006	716	0.48	-2.18

Source: GSOEP 1995, own calculations.

Notes: 1) Coefficient of the schooling variable α_1. To obtain the return in percent, compute $[\exp(\alpha_1)-1] \times 100$.
2) H_0: $\text{Return}_{\text{model i}} = \text{Return}_{\text{reference model}}$. t-statistics $= (b_2 - b_1) / \sqrt{(se_1)^2 + (se_2)^2}$. If $|t|>1.96$ (resp. 2.58, 1.65), then the hypothesis of equality of the coefficients is rejected at a significance level of 5% (resp. 1%, 10%).

Unless age is used as a proxy for labour market experience, male returns to schooling prove quite robust across the different specifications. When age is used as a proxy, male returns to schooling are lower than for any other specification. For women, using age also reduces the return to schooling in a significant way. Using actual experience (total or full-time only) instead of potential experience does not change the results for men, and only slightly – but not significantly - for women. The same holds for model 4, which includes previous non-employment. Adding previous unemployment duration reduces the returns to schooling, but not dramatically; the same goes for the inclusion of tenure and the part-time dummy. However, the inclusion of a very large number of control variables for firm size, industry and region reduces the return to schooling significantly for women, so that gender differences are very small under this specification.

On the whole, male returns to schooling are not sensitive to the specification of the wage equation. Female returns are a little bit more sensitive, especially with respect to controls for previous non-employment, as well as for industry, firm size and region. Adding these variables also improves significantly the fit of the regression.

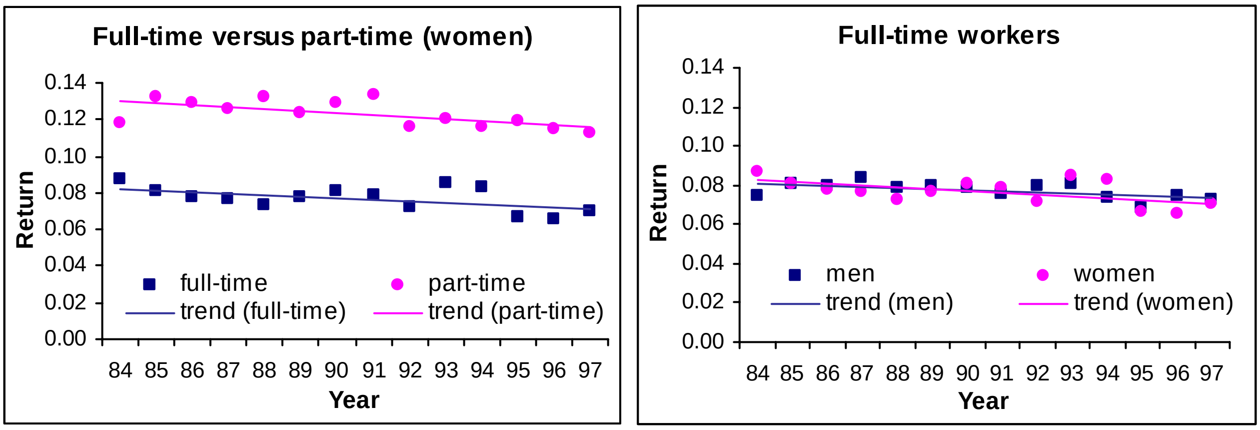
3.3 The choice of the sample

The general statement drawn from the analysis of the standard model may hide some more complex developments. In this section, we focus on two issues: differences between full-time and part-time workers and between private and public sectors.

3.3.1 Full-time versus part-time workers

Although we used in our basic model gross hourly wage as the dependent variable and not monthly wage, for instance, the estimated returns to education may be affected by differences in working time. Part-time employment is a virtually non-existent phenomenon among men in Germany, but about 40% of our female sample work only part-time. Figure 7 represents the returns to education for part-time working women compared to those working full-time (left hand-side) and gender differences among full-timers only (right hand-side).

Figure 7: Return[1)] to education 1984-97 - Full-timers versus part-timers



Source: GSOEP 1984-97, own calculations.
Note: 1) Coefficient of the schooling variable α_1. To obtain the return in percent, compute $[\exp(\alpha_1)\text{-}1] \times 100$.

The results are striking. Overall, we found a return of about 10% for women in the reference model. However, there are huge differences between full-timers and part-timers. Part-timers have a much higher return to education (around 13%) than full-time working women (around 8%). Hence, wage discrepancies between higher and lower educated workers seem larger among part-time workers. Furthermore, the returns to education for full-time workers are similar for men and for women. The slight downward trend seems comparable between part-time and full-time working women.

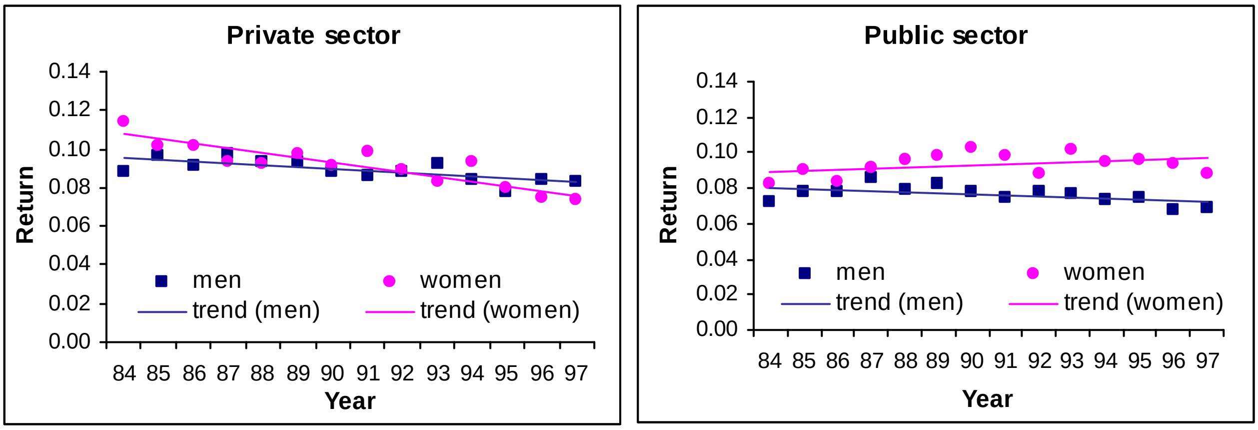
3.3.2 Private versus public sector

Wage setting mechanisms are not the same in the public and in the private sector[8]. Thus, the wage structure is expected to differ and so are returns to education. Figure 8 shows the estimated returns to education in the private and in the public sector of the economy. Again, the differences are remarkable. In the public sector, female returns to education tend to increase over the period, contrary to male returns, which decrease. As a result, gender differences in the return to education have increased in favour of women. In the private sector, no such developments are observable. Female returns to education have decreased significantly and more sharply than those for men. As a result, although female returns in the

[8] For instance, wages are indexed on age in the public sector.

private sector were higher than for men at the beginning of the 1980s, they have become lower in the 1990s.

Figure 8: Return[1)] to education 1984-97 - Private versus public sector



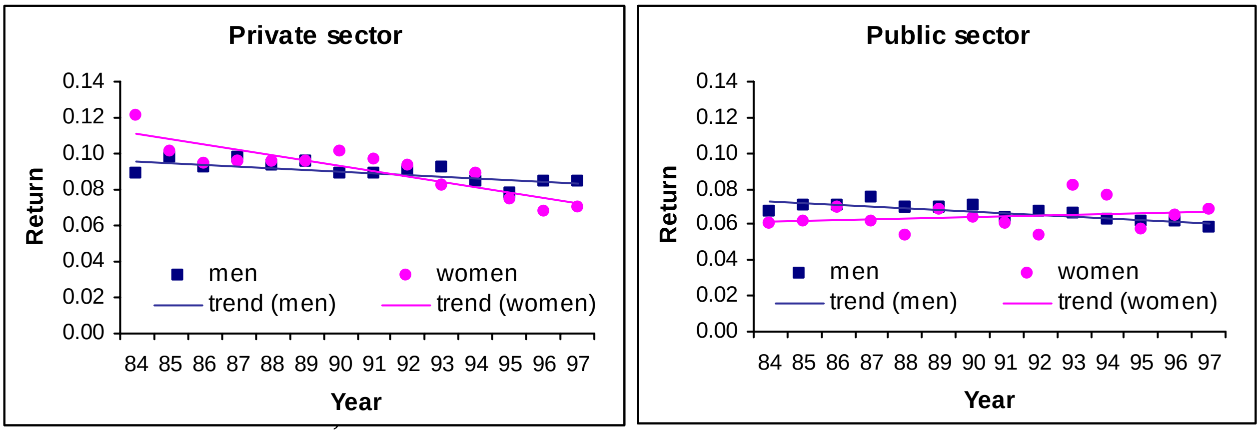
Source: GSOEP 1984-97, own calculations.

Note: 1) Coefficient of the schooling variable α_1. To obtain the return in percent, compute [exp(α_1)-1] ×100.

3.3.3 Full-time employees: private versus public sector

The preceding analysis has revealed marked differences between full-time and part-time workers, on the one hand, and between private and public sectors on the other hand. Now, we combine these effects and examine returns to education of full-time employees in the two sectors of the economy.

Figure 9: Return[1)] to education 1984-97 - Private versus public sector, full-time workers



Note: 1) Coefficient of the schooling variable α_1. To obtain the return in percent, compute [exp(α_1)-1] ×100.

Figure 9 shows that the downward trend in female returns to education in the private sector is more pronounced if only full-time employees are taken into account (from 0.12 to less than 0.07). In the public sector, the level of the returns is significantly lower than in the case where all employees are considered, whether they work full-time or part-time. However, the upward trend in the returns to education of female full-time employees is more pronounced than that of employees overall.

On the whole, the choice of the sample of observation plays a crucial role concerning both the trend and the level of the returns to education. Therefore, returns to education can only be compared if the same sample is analysed.

3.4 Choice of the estimation method

So far, we have estimated our wage equations with Ordinary Least Squares (OLS). However, much of the recent literature focuses on the estimation biases which may arise while estimating empirical wage equations (see Card 1999). In particular, the issue of sample selectivity and the endogeneity of schooling are often mentioned as possible sources of biases. This section concentrates on these two issues.[9]

3.4.1 Correcting for the selectivity bias of participation in employment

Selectivity bias occurs if the expectation of the dependent variable, given the set of exogenous explanatory variables, differs from its expectation given these control variables and some other conditioning choice variable (see Heckman 1979). In the context of the estimation of wage functions, the individual decision to work will determine whether we observe the person's wage in our data. If the factors determining this decision are uncorrelated with the factors affecting individual wages, we could simply ignore the fact that not all wages are observed. However, such an assumption is unlikely to hold in practice, especially for women, because women with higher market wages probably tend to participate more in the labour force. Hence, employed women are likely to be a self-selected group whose wages may not be representative of all women with given observed characteristics, which could bias estimated returns to education.

- Sample selection into non-employment and employment

We apply a full maximum-likelihood procedure to correct for potential sample selection bias, i.e., we estimate the wage and the participation equations simultaneously. This procedure requires the availability of some credible instruments, i.e. variables significantly affecting labour force participation but having no significant direct effect on earnings. We test whether the standard specification is robust towards correction for selectivity bias, and whether the choice of the instruments matters. We estimate the following models[10]:

- Reference model: No correction for selectivity bias.
- Model 1: We use marital status and the number of children below the age of 16 years as instruments for the selection equation.
- Model 2: We use household financial variables as exclusion restrictions (other net household income and the amount of monthly mortgage payments).
- Model 3: We use variables indicating whether mother and father were predominantly employed during childhood, and whether the person grew in a rural area, medium city or large city.
- Model 4: We use both marital status/number of children and household financial variables.
- Model 5: We include all variables together as exclusion restrictions.

In Table 7 and Table 8, we report the estimation results of the different models with respect to the return to schooling, the coefficient of the selectivity correction term λ and their

[9] Of course, there are other possible sources of biases, like omitted variable bias (in particular ability) or measurement error. Since we do not have information to account for these factors, there is little we can do here.

[10] All models (except the reference model) entail the variables of the main equation in the selection equation (years of schooling, potential experience and potential experience squared).

respective standard errors. Additionally, we run two tests: a test of collinearity between the inverse Mill's ratio and the regressors of the regression (see Puhani 1997), and a t-test to check whether the corrected coefficients of the various models significantly differ from the standard estimates without correction.

Table 7: Sensitivity of the returns[1)] to education to selectivity bias correction (1995) - Men

	Estimation results					Tests	
	Return	Std. err.	λ	Std. err.	Obs.	Collinearity[2)]	t-test[3)]
Ref. model	0.073	0.003	-	-	1138	-	-
Model 1	0.074	0.003	-0.196	0.030	1138	0.14	0.07
Model 2	0.071	0.003	-0.189	0.021	1119	0.25	-0.60
Model 3	0.076	0.004	-0.119	0.142	868	0.84	0.49
Model 4	0.071	0.003	-0.171	0.021	1119	0.06	-0.50
Model 5	0.074	0.004	-0.170	0.023	851	0.08	0.07

Source: GSOEP 1995, own calculations.

Notes: 1) Coefficient of the schooling variable α_1. To obtain the return in percent, compute $[\exp(\alpha_1)-1] \times 100$. 2) R^2 of the regression of the inverse Mill's ratio on the regressors of the main equation. 3) H_0: $\text{Return}_{\text{model i}} = \text{Return}_{\text{reference model}}$. t-statistics $= (b_2 - b_1) / \sqrt{(se_1)^2 + (se_2)^2}$. If $|t|>1.96$ (resp. 2.58, 1.65), then the hypothesis of equality of the coefficients is rejected at a significance level of 5% (resp. 1%, 10%).

Table 8: Sensitivity of the returns[1)] to education to selectivity bias correction (1995) - Women

	Estimation results					Tests	
	Return	Std. err.	λ	Std. err.	Obs.	Collinearity[2)]	t-test[3)]
Ref. model	0.097	0.006	-	-	722	-	-
Model 1	0.081	0.007	-0.538	0.040	722	0.73	-1.70
Model 2	0.078	0,017	-0.439	0.034	705	0.43	-1.00
Model 3	0.087	0.008	-0.535	0.044	562	0.96	-0.91
Model 4	0.083	0.007	-0.403	0.037	705	0.25	-1.54
Model 5	0.085	0.007	-0.410	0.042	547	0.26	-1.21

Source: GSOEP 1995, own calculations.

Notes: 1) Coefficient of the schooling variable α_1. To obtain the return in percent, compute $[\exp(\alpha_1)-1] \times 100$. 2) R^2 of the regression of the inverse Mill's ratio on the regressors of the main equation. 3) H_0: $\text{Return}_{\text{model i}} = \text{Return}_{\text{reference model}}$. t-statistics $= (b_2 - b_1) / \sqrt{(se_1)^2 + (se_2)^2}$. If $|t|>1.96$ (resp. 2.58, 1.65), then the hypothesis of equality of the coefficients is rejected at a significance level of 5% (resp. 1%, 10%).

Estimation results[11] from the selection equation show that most of the variables chosen as potential instruments are highly significant and have a strong effect on labour-force participation. Being married or having children in the household, for instance, strongly reduces the probability of employment for women, as opposed to the fact that the parents were predominantly employed during the individual's childhood. For men, however, being married increases the probability of being employed, whereas the employment situation of the parents during childhood has no significant effect. Financial variables have a significant - though small - influence, since the higher other household income, the smaller the probability of being employed. Conversely, individuals in more highly indebted households are more likely to participate in the labour market.

For men, except for model 3, all estimates yield significant negative coefficients of the selectivity-correction term λ, which suggests the presence of some selectivity bias. The value of the coefficient varies from –0.17 to –0.19. Nevertheless, the returns to schooling are not

[11] Available on request.

affected much by selectivity correction: the "corrected" coefficients are very close to the OLS estimates and the difference is negligible. For women, all selectivity correction terms λ are negative and highly significant, ranging from -0,40 to -0,54. All models yield selectivity-corrected returns to schooling which are below the OLS return. However, the difference between the selectivity-corrected coefficient and the OLS coefficient is only significant (at the 10% level) in model 1, using marital status and the number of children as sole exclusion restrictions. For model 3 the collinearity test indicates that the instruments are rather weak. In model 5, where we also use marital status and number of children as exclusion in addition to other variables, the coefficient on λ is not significant. Therefore, the size and the significance of the selectivity correction term seems somewhat sensitive to the choice of instruments and this correction should therefore be used with care. Moreover, even if the λ turned out to be significant in most cases, overall we find no evidence for a sample selectivity bias with respect to the returns to schooling for men, and only rather small effects for women.

- Sample selection into non-employment, part-time and full-time employment

The preceding correction for selectivity bias makes no distinction whether women participate in full-time or in part-time employment. However, as we saw, the return to education is higher for part-time working women. Thus, it may be of interest to examine whether the coefficients estimated by OLS for women working part-time compared to full-time female employees are biased. Here, we adapt the two-step Heckman (1979) procedure and first estimate the selection equation by an ordered probit model, where the dependent variable is the probability of being either not employed, part-time employed, or full-time employed. This enables us to compute two selectivity correction terms, for part-time employment and for full-time employment, which we include as additional regressors in the second-step wage equations. We keep the five alternative models defined above, using the same exclusion restrictions, and compare them with OLS estimations for part-time and full-time workers, respectively.

As Table 9 and Table 10 show, although the selectivity correction terms are negative and significant for both part-time and full-time workers (except for model 3), there is not much difference in the coefficients whether one corrects for sample selectivity or not. The corrected coefficients have the same order of magnitude as the OLS estimates, though the standard errors are relatively large.

Table 9: Sensitivity of the returns[1)] to education to selectivity bias correction by ordered probit (1995) – Part-time working women

	Estimation results					Tests	
	Return	Std. err.	λ	Std. err.	Obs.	Collinearity[2)]	t-test[3)]
Ref. model	0.119	0.008	-	-	378	-	-
Model 1	0.115	0.010	-0.116	0.010	378	0.42	-0.34
Model 2	0.106	0.009	-0.253	0.072	365	0.31	-1.08
Model 3	0.115	0.018	-0.485	0.694	312	0.95	-0.20
Model 4	0.113	0.009	-0.113	0.043	365	0.13	-0.53
Model 5	0.120	0.010	-0.105	0.052	301	0.15	0.09

Source: GSOEP 1995, own calculations.

Notes: 1) Coefficient of the schooling variable α_1. To obtain the return in percent, compute $[\exp(\alpha_1)-1] \times 100$. 2) R² of the regression of the inverse Mill's ratio on the regressors of the main equation. 3) H_0: $\text{Return}_{\text{model i}} = \text{Return}_{\text{reference model}}$. t-statistics = $(b_2 - b_1) / \sqrt{(se_1)^2 + (se_2)^2}$. If $|t|$>1.96 (resp. 2.58, 1.65), then the hypothesis of equality of the coefficients is rejected at a significance level of 5% (resp. 1%, 10%).

Table 10: Sensitivity of the returns[1)] to education to selectivity bias correction by ordered probit (1995) – Full-time working women

	Estimation results					Tests	
	Return	Std. err.	λ	Std. err.	Obs.	Collinearity[2)]	t-test[3)]
Ref. model	0.066	0.008	-	-	344	-	-
Model 1	0.058	0.009	-0.353	0.149	344	0.54	-0.73
Model 2	0.064	0.008	-0.113	0.055	340	0.05	-0.21
Model 3	0.070	0.016	0.281	0.709	250	0.94	0.22
Model 4	0.065	0.008	-0.162	0.045	340	0.15	-0.11
Model 5	0.063	0.009	-0.183	0.049	246	0.12	-0.22

Source: GSOEP 1995, own calculations.

Notes: 1) Coefficient of the schooling variable α_1. To obtain the return in percent, compute [exp(α_1)-1] ×100. 2) R^2 of the regression of the inverse Mill's ratio on the regressors of the main equation. 3) H_0: $\text{Return}_{\text{model i}}$ = $\text{Return}_{\text{reference model}}$. t-statistics = $(b_2 - b_1) / \sqrt{(se_1)^2 + (se_2)^2}$. If $|t|$>1.96 (resp. 2.58, 1.65), then the hypothesis of equality of the coefficients is rejected at a significance level of 5% (resp. 1%, 10%).

3.4.2 Accounting for the endogeneity of schooling

Until now, we supposed that human capital variables are exogenous. Obviously, this may not be the case. Here, we examine the effect of allowing the schooling variable to be endogenous. This requires the availability of variables which affect educational attainment and, at the same time, have no direct effect on wages. Unfortunately, we do not dispose of indicators for learning ability, intelligence or motivation, which are undoubtedly essential factors affecting performance at school. However, the GSOEP does provide information about family background, which is likely to affect educational attainment. Here, we follow a two-stage instrumental variables (IV) procedure, where the schooling variable is first regressed on a set of explanatory variables and then instrumented by its predicted value in the second-stage wage equation.

Again, we test for alternative models using different instruments in order to inspect both the sensitivity of the return to education to the correction for endogeneity of schooling and to the choice of the instruments. We estimate the following models:

- Reference model: OLS estimation (i.e. schooling assumed exogenous).
- Model 1: The schooling variable is instrumented by the level of education of the mother and of the father, expressed in years of schooling.
- Model 2: The occupational position of the father when the individual was 15 is used as an instrument (not-employed/blue collar versus white collar, self-employed or civil servant).
- Model 3: Both the educational degree and the occupation of the father are used.
- Model 4: Dummies indicating whether the parents were predominantly employed during the individual's childhood, whether the person grew up with both parents, and whether the family lives in a rural or urban community serve as instruments.
- Model 5: All variables together.

Most of the variables chosen as instruments have a significant and strong effect on the level of educational attainment[12]. The better educated the parents are, the higher the own educational level tends to be. The educational level of the father seems to play a more important role than that of the mother, especially for women. The occupational position of the father also seems to be of crucial importance. Sons and daughters of civil servants have the best chances to become highly educated, followed by children of white-collar workers. For both genders, individuals whose parents are blue-collar workers or not employed have the

[12] Results are available on request.

worst educational prospects. Whether the person grew up with both parents or not seems to have a positive impact on males' educational attainment but no relevance for women. Altogether, these variables explain about 20% of the variance of schooling.

Table 11: Sensitivity of returns[1)] to education: OLS versus IV with endogenous schooling (1995)

	Men					Women				
	Return	Std. err.	Obs.	R^2	t-test[2)]	Return	Std. err.	Obs.	R^2	t-test[2)]
Ref. model	0.076	0.003	1464	0.34		0.097	0.006	722	0.32	
Model 1	0.080	0.008	1026	0.10	0.38	0.088	0.014	675	0.12	-0.60
Model 2	0.066	0.011	1138	0.04	-0.91	0.103	0.016	722	0.11	0.37
Model 3	0.077	0.008	1026	0.10	0.12	0.093	0.013	675	0.14	-0.30
Model 4	0.148	0.046	868	0.02	1.54	0.112	0.060	562	0.08	0.26
Model 5	0.075	0.008	815	0.10	-0.17	0.106	0.014	536	0.16	0.60

Source: GSOEP 1995, own calculations.

Notes: 1) Coefficient of the schooling variable α_1. To obtain the return in percent, compute $[\exp(\alpha_1)-1] \times 100$. 2) R^2 of the regression of the inverse Mill's ratio on the regressors of the main equation. 3) H_0: $\text{Return}_{\text{model i}} = \text{Return}_{\text{reference model}}$. t-statistics $= (b_2 - b_1) / \sqrt{(se_1)^2 + (se_2)^2}$. If $|t|>1.96$ (resp. 2.58, 1.65), then the hypothesis of equality of the coefficients is rejected at a significance level of 5% (resp. 1%, 10%).

The results are somewhat sensitive to the choice of the instruments. However, the differences between the OLS and the instrumented returns to schooling are statistically insignificant in all models. This is sometimes due to the fact that the coefficients have the same order of magnitude (e.g. model 3 for men), or that the standard error is very large (e.g. model 4 for men or model 1 for women). Thus, no clear conclusion can be drawn from these estimations regarding the existence and even the direction of a supposed endogeneity bias. It is imperative to have better instruments at one's disposal if one intends to correct for the endogeneity of schooling.

4 Conclusions

Our empirical analysis has provided a broad assessment of the returns to education in West Germany over the past decades. In a first step, simple estimates of the returns to schooling based on standard Mincer equations showed that the returns to one additional year of schooling have remained remarkably stable since the mid-1980s. Women have significantly higher returns to schooling (about 10%) than men (about 8%). In a second step, we tested this specification in various ways, taking into account possible cohort effects, the choice of the sample, the definition of the human capital variables and different estimation methods.

Firstly, the analysis showed that the developments in the returns to education result from cohort and life-cycle effects in addition to time-effects. We found evidence that the returns to education are not constant over the life-cycle, especially for women. Evaluating the returns to schooling for different cohorts at the same age shows that a significant decline in the returns to education across cohorts is observable at age 30 to 39, and this decline is particularly pronounced for women since 1994. At the middle of the career (age 40-49), we found evidence for slightly increasing (men) or constant (women) returns across birth cohorts. Finally, at an older age (50 to 60), the returns to education are lower for younger cohorts, particularly for women beginning in 1994.

Secondly, we examined whether differences in the specification of human capital variables in the wage equation could alter the estimates of the returns to education. Departing from the quantitative measure of education, we analysed the returns to educational degrees. The higher the degree, the higher the wage premium. However, when we correct for the different length of studies associated with the various degrees, we find that the master degree yields the highest returns. A downward trend is observable for the return to the master degree and to the high school diploma at the end of the period. Our tests for alternative specifications of the other variables showed that the level of the return to schooling is quite robust. Only when we use age instead of labour market experience, the returns to education are somewhat lower. Female returns to education are somewhat sensitive to the inclusion of additional variables designed to capture previous non-employment or to control for the industrial and regional structure, but the difference from the simple specification is not really important.

Thirdly, the analysis reveals huge differences across subgroups of workers. The returns to education are much higher for part-time than for full-time working women (12-13% versus 8%). Taking full-time workers only, there are no significant gender differences: the return amounts to about 8% for both genders. Moreover, the differences between public and private sectors are also remarkable. In the public sector, female returns to schooling have increased somewhat (from about 9% to 10%) over the period, whereas male returns have slightly decreased (from about 8% to 7%). As a result, female returns to schooling are significantly higher than male returns. In the private sector, however, female returns decrease much stronger than male ones. Consequently, women have now lower returns to schooling than men. This trend is even more pronounced if one only considers full-time employees in the private sector.

Finally, we examined methodological issues and focussed on the impact of possible selectivity and endogeneity biases on the level of the returns to education. The results were somewhat inconclusive and do not point to significant estimation biases. Whereas the selectivity correction term proved mostly significant (negative for both gender, and stronger for women), which points to the presence of selectivity bias, the returns to schooling were not affected by the selectivity correction in a significant way. The same results hold when we distinguish between selection into part-time and full-time employment. We adopted a similar approach to check for the endogeneity of schooling. Again, the results proved somewhat sensitive to the choice of the instruments to explain educational attainment and no conclusion can be drawn as to the size or even the direction of the supposed endogeneity bias.

On the whole, the simple estimates proved quite robust towards specification and estimation method. However, the overall assessment hides some more complex developments, i.e. huge differences between subgroups of workers or conjunction of time, life-cycle and cohort effects. Thus, studies on the returns to education in West Germany should be interpreted very carefully and one should be aware of the implications of the specific framework adopted.

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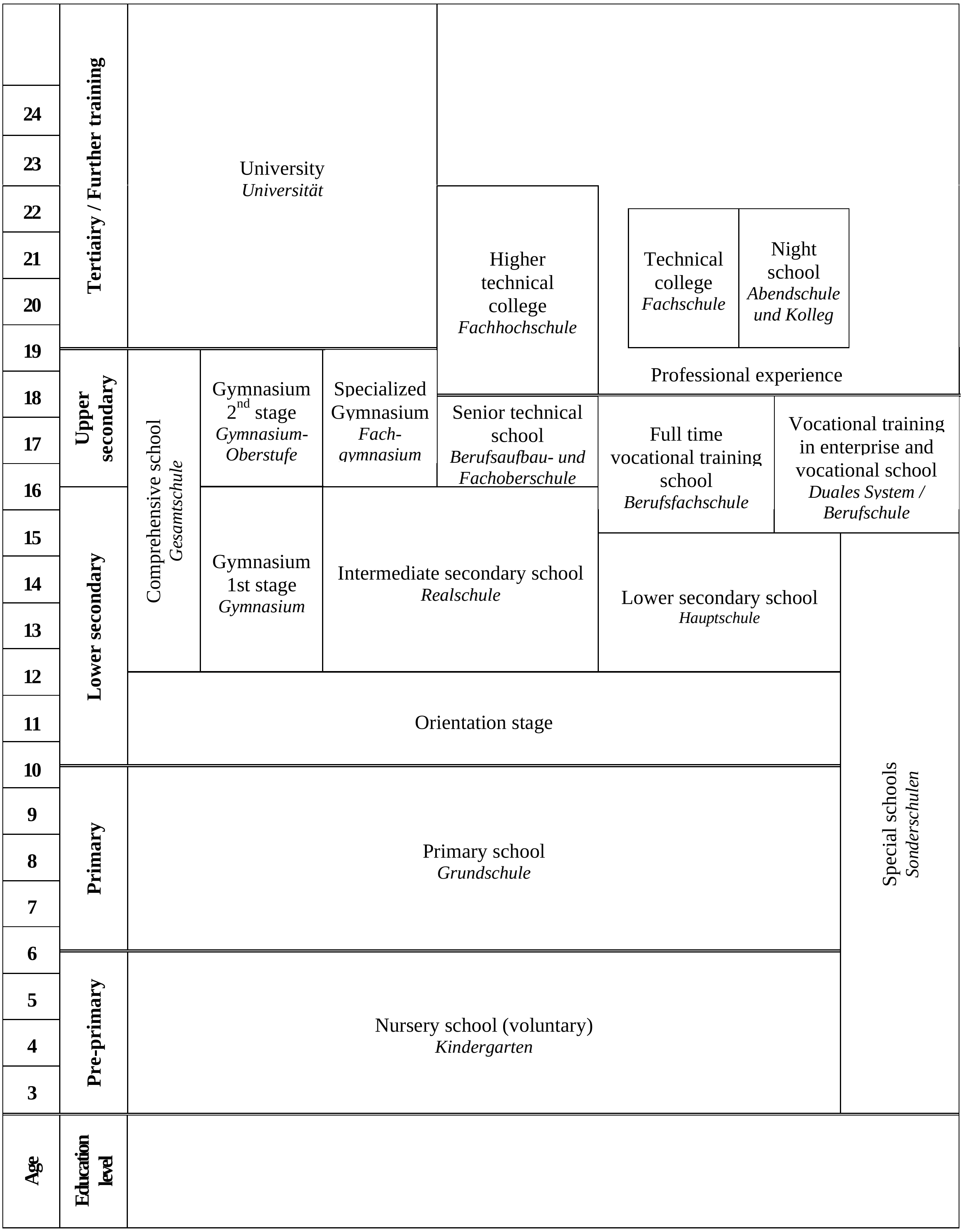
Annex

Annex 1: The German education system



Annex 2: Return[1)] to education 1984-97 - Dependent. variable: log gross hourly wage

	Men				Women				t-test
	Return	Std. err.	Obs.	R^2	Return	Std. err.	Obs.	R^2	Men=Women
1984	0.076	0.003	1464	0.34	0.103	0.007	785	0.29	-3.73
1985	0.084	0.003	1340	0.34	0.105	0.006	758	0.32	-3.05
1986	0.082	0.003	1303	0.35	0.101	0.006	733	0.29	-2.64
1987	0.088	0.003	1296	0.40	0.102	0.006	733	0.32	-2.03
1988	0.082	0.003	1287	0.37	0.103	0.006	712	0.31	-3.03
1989	0.084	0.003	1221	0.39	0.100	0.006	693	0.30	-2.31
1990	0.080	0.003	1186	0.37	0.105	0.006	674	0.31	-3.46
1991	0.078	0.003	1138	0.37	0.108	0.006	715	0.33	-4.45
1992	0.082	0.003	1118	0.40	0.097	0.006	739	0.33	-2.26
1993	0.084	0.003	1158	0.39	0.103	0.006	725	0.31	-2.77
1994	0.077	0.003	1121	0.37	0.100	0.006	736	0.33	-3.57
1995	0.073	0.003	1138	0.33	0.097	0.006	722	0.32	-3.49
1996	0.076	0.003	1120	0.33	0.092	0.006	769	0.27	-2.48
1997	0.074	0.003	1098	0.33	0.091	0.005	746	0.29	-2.72
Pooled	0.080	0.001	16988	0.39	0.100	0.002	10240	0.34	-11.18

Source: GSOEP 1995, own calculations.

Notes: 1) Coefficient of the schooling variable α_1. To obtain the return in percent, compute $[\exp(\alpha_1)-1] \times 100$.
2) H_0: $\text{Return}_{men} = \text{Return}_{women}$. t-statistics = $(b_2 - b_1) / \sqrt{(se_1)^2 + (se_2)^2}$. If $|t|>1.96$ (resp. 2.58, 1.65), then the hypothesis of equality of the coefficients is rejected at a significance level of 5% (resp. 1%, 10%).

Annex 3: Return[1)] to education 1984-97 - Full-time versus part-time working women

	Full-time				Part-time				t-test
	Return	Std. err.	Obs.	R^2	Return	Std. err.	Obs.	R^2	FT=PT
1984	0.087	0.007	374	0.34	0.118	0.011	411	0.27	-2.40
1985	0.080	0.007	385	0.25	0.132	0.010	373	0.41	-4.26
1986	0.078	0.007	393	0.24	0.130	0.011	340	0.38	-3.97
1987	0.077	0.007	383	0.25	0.125	0.009	350	0.41	-4.13
1988	0.073	0.007	369	0.25	0.132	0.010	343	0.40	-4.87
1989	0.077	0.007	367	0.25	0.123	0.010	326	0.36	-3.74
1990	0.080	0.008	348	0.25	0.129	0.010	326	0.39	-3.86
1991	0.079	0.008	345	0.26	0.133	0.009	370	0.41	-4.63
1992	0.072	0.008	352	0.23	0.116	0.008	387	0.42	-4.06
1993	0.085	0.008	364	0.25	0.120	0.009	361	0.38	-2.91
1994	0.083	0.007	362	0.28	0.115	0.008	374	0.40	-2.98
1995	0.066	0.008	344	0.20	0.119	0.008	378	0.41	-4.60
1996	0.066	0.008	377	0.17	0.115	0.008	392	0.37	-4.37
1997	0.070	0.007	363	0.22	0.112	0.008	383	0.38	-4.05

Source: GSOEP 1995, own calculations.

Notes: 1) Coefficient of the schooling variable α_1. To obtain the return in percent, compute $[\exp(\alpha_1)-1] \times 100$.
2) H_0: $\text{Return}_{full-time} = \text{Return}_{part-time}$. t-statistics = $(b_2 - b_1) / \sqrt{(se_1)^2 + (se_2)^2}$. If $|t|>1.96$ (resp. 2.58, 1.65), then the hypothesis of equality of the coefficients is rejected at a significance level of 5% (resp. 1%, 10%).

Annex 4: Return[1)] to education 1984-97 by age groups - Men

	Age group 30-39				Age group 40-49				Age group 50-60			
	return	s.e.	obs.	R^2	return	s.e.	obs.	R^2	return	s.e.	obs.	R^2
1984	0.080	0.007	458	0.02	0.076	0.006	608	0.39	0.081	0.008	382	0.38
1985	0.081	0.008	451	0.25	0.082	0.007	519	0.42	0.083	0.010	354	0.33
1986	0.079	0.006	463	0.33	0.093	0.008	475	0.38	0.084	0.010	345	0.32
1987	0.097	0.006	468	0.39	0.095	0.007	437	0.39	0.086	0.009	367	0.39
1988	0.090	0.006	464	0.36	0.086	0.007	418	0.34	0.093	0.010	376	0.38
1989	0.078	0.006	446	0.31	0.083	0.007	400	0.35	0.100	0.010	352	0.46
1990	0.064	0.007	451	0.27	0.077	0.008	368	0.34	0.090	0.010	352	0.46
1991	0.071	0.006	423	0.30	0.086	0.008	357	0.34	0.093	0.009	340	0.44
1992	0.079	0.006	442	0.34	0.097	0.009	334	0.37	0.090	0.009	331	0.49
1993	0.080	0.006	462	0.34	0.095	0.009	331	0.36	0.086	0.009	353	0.47
1994	0.074	0.006	467	0.32	0.089	0.009	317	0.39	0.085	0.009	325	0.43
1995	0.065	0.006	471	0.21	0.089	0.008	357	0.36	0.083	0.009	293	0.46
1996	0.068	0.007	478	0.21	0.086	0.008	363	0.37	0.089	0.010	263	0.44
1997	0.065	0.007	477	0.18	0.094	0.008	369	0.38	0.082	0.009	239	0.44

Source: GSOEP 1984-97, own calculations.

Note: 1) Coefficient of the schooling variable α_1. To obtain the return in percent, compute $[\exp(\alpha_1)-1] \times 100$.

Annex 5: Return[1)] to education 1984-97 by age groups - Women

	Age group 30-39				Age group 40-49				Age group 50-60			
	return	s.e.	obs.	R^2	return	s.e.	obs.	R^2	return	s.e.	obs.	R^2
1984	0.086	0.014	281	0.29	0.094	0.016	331	0.23	0.126	0.018	168	0.40
1985	0.104	0.012	288	0.33	0.119	0.017	286	0.26	0.130	0.016	176	0.40
1986	0.090	0.013	274	0.26	0.111	0.016	281	0.33	0.126	0.019	172	0.30
1987	0.085	0.009	271	0.31	0.107	0.017	275	0.31	0.130	0.014	178	0.42
1988	0.093	0.011	275	0.33	0.113	0.017	243	0.32	0.127	0.016	186	0.34
1989	0.098	0.011	268	0.36	0.104	0.018	233	0.25	0.123	0.016	184	0.35
1990	0.101	0.011	272	0.38	0.106	0.018	218	0.22	0.128	0.015	178	0.37
1991	0.099	0.010	273	0.33	0.115	0.016	235	0.29	0.139	0.016	201	0.37
1992	0.088	0.010	290	0.32	0.112	0.015	242	0.36	0.136	0.018	200	0.28
1993	0.096	0.011	282	0.31	0.109	0.016	255	0.32	0.153	0.026	182	0.28
1994	0.094	0.010	295	0.32	0.119	0.015	249	0.34	0.147	0.020	188	0.34
1995	0.078	0.011	298	0.21	0.116	0.014	244	0.41	0.136	0.020	170	0.32
1996	0.072	0.010	327	0.20	0.110	0.013	267	0.29	0.127	0.018	164	0.44
1997	0.064	0.010	318	0.21	0.106	0.012	257	0.37	0.107	0.020	159	0.35

Source: GSOEP 1984-97, own calculations.

Notes: 1) Coefficient of the schooling variable α_1. To obtain the return in percent, compute $[\exp(\alpha_1)-1] \times 100$.

Annex 6: Return[1)] to education 1984-97 - Full-time working men versus full-time working women

	Men				Women				t-test[2)]
	Return	Std. err.	Obs.	R^2	Return	Std. err.	Obs.	R^2	Men=Women
1984	0.075	0.003	1444	0.32	0.087	0.007	374	0.34	-1.61
1985	0.081	0.003	1317	0.33	0.080	0.007	385	0.25	0.08
1986	0.079	0.003	1276	0.32	0.078	0.007	393	0.24	0.20
1987	0.084	0.003	1265	0.35	0.077	0.007	383	0.25	0.92
1988	0.078	0.003	1259	0.34	0.073	0.007	369	0.25	0.72
1989	0.079	0.003	1194	0.34	0.077	0.007	367	0.25	0.29
1990	0.078	0.003	1157	0.34	0.080	0.008	348	0.25	-0.26
1991	0.076	0.003	1102	0.35	0.079	0.008	345	0.26	-0.40
1992	0.080	0.003	1086	0.37	0.072	0.008	352	0.23	0.97
1993	0.081	0.003	1125	0.35	0.085	0.008	364	0.25	-0.45
1994	0.073	0.003	1087	0.35	0.083	0.007	362	0.28	-1.17
1995	0.068	0.003	1093	0.30	0.066	0.008	344	0.20	0.27
1996	0.075	0.003	1081	0.32	0.066	0.008	377	0.17	1.05
1997	0.072	0.003	1047	0.31	0.070	0.007	363	0.22	0.26

Source: GSOEP 1984-97, own calculations.

Notes: 1) Coefficient of the schooling variable α_1. To obtain the return in percent, compute $[\exp(\alpha_1)-1] \times 100$.
2) H_0: $Return_{men} = Return_{women}$. t-statistics = $(b_2 - b_1) / \sqrt{(se_1)^2 + (se_2)^2}$. If $|t|>1.96$ (resp. 2.58, 1.65), then the hypothesis of equality of the coefficients is rejected at a significance level of 5% (resp. 1%, 10%).

Annex 7: Return[1)] to education 1984-97 - Private versus public sector - Men

	Private				Public				t-test[2)]
	Return	Std. err.	Obs.	R^2	Return	Std. err.	Obs.	R^2	Private=Public
1984	0.088	0.004	985	0.31	0.072	0.004	475	0.44	2.69
1985	0.097	0.005	883	0.31	0.078	0.004	436	0.45	2.95
1986	0.092	0.005	837	0.31	0.078	0.004	415	0.49	2.11
1987	0.098	0.005	824	0.33	0.086	0.004	415	0.59	1.85
1988	0.093	0.005	821	0.33	0.079	0.004	404	0.56	2.44
1989	0.094	0.004	841	0.35	0.083	0.004	374	0.53	1.76
1990	0.088	0.004	823	0.33	0.078	0.004	348	0.50	1.58
1991	0.086	0.004	783	0.34	0.074	0.004	346	0.50	1.98
1992	0.088	0.004	768	0.38	0.078	0.004	332	0.51	1.69
1993	0.093	0.004	798	0.35	0.077	0.004	347	0.52	2.52
1994	0.084	0.004	746	0.33	0.074	0.004	334	0.49	1.75
1995	0.078	0.004	789	0.29	0.075	0.004	341	0.50	0.50
1996	0.084	0.004	773	0.33	0.068	0.004	311	0.52	2.67
1997	0.083	0.004	780	0.32	0.068	0.005	317	0.44	2.33

Source: GSOEP 1984-97, own calculations.

Notes: 1) Coefficient of the schooling variable α_1. To obtain the return in percent, compute $[\exp(\alpha_1)-1] \times 100$.
2) H_0: $Return_{private} = Return_{public}$. t-statistics = $(b_2 - b_1) / \sqrt{(se_1)^2 + (se_2)^2}$. If $|t|>1.96$ (resp. 2.58, 1.65), then the hypothesis of equality of the coefficients is rejected at a significance level of 5% (resp. 1%, 10%).

Annex 8: Return[1)] to education 1984-97 - Private versus public sector - Women

	Private				**Public**				**t-test[2)]**
	Return	Std. err.	Obs.	R²	Return	Std. err.	Obs.	R²	Private=Public
1984	0.115	0.006	513	0.16	0.082	0.006	269	0.47	3.61
1985	0.102	0.006	435	0.16	0.090	0.006	272	0.50	1.41
1986	0.102	0.006	395	0.14	0.084	0.006	258	0.48	2.06
1987	0.094	0.006	389	0.15	0.091	0.006	259	0.52	0.27
1988	0.092	0.007	386	0.16	0.096	0.007	259	0.49	-0.47
1989	0.097	0.007	412	0.17	0.098	0.007	275	0.42	-0.06
1990	0.091	0.007	395	0.16	0.103	0.007	257	0.49	-1.17
1991	0.099	0.006	445	0.17	0.098	0.006	261	0.58	0.04
1992	0.089	0.006	441	0.17	0.088	0.006	278	0.51	0.14
1993	0.083	0.007	446	0.14	0.101	0.007	271	0.49	-1.96
1994	0.094	0.006	435	0.21	0.095	0.007	266	0.47	-0.13
1995	0.080	0.007	442	0.17	0.095	0.007	272	0.47	-1.68
1996	0.075	0.006	448	0.12	0.094	0.006	277	0.48	-2.16
1997	0.073	0.005	462	0.12	0.088	0.005	280	0.51	-1.95

Source: GSOEP 1984-97, own calculations.

Notes: 1) Coefficient of the schooling variable α_1. To obtain the return in percent, compute [exp(α_1)-1] ×100.
2) H_0: $Return_{private} = Return_{public}$. t-statistics = $(b_2 - b_1) / \sqrt{(se_1)^2 + (se_2)^2}$. If $|t|$>1.96 (resp. 2.58, 1.65), then the hypothesis of equality of the coefficients is rejected at a significance level of 5% (resp. 1%, 10%).

Annex 9: Wage premia[1)] associated with education degrees 1984-97 - Men

	Apprentice.		Master		High school		High. tech. coll.		University		Obs.	R²
	return	s.e.	return	s.e.	return	s.e.	return	s.e.	return	s.e.		
1984	0.128	0.023	0.256	0.027	0.438	0.039	0.509	0.037	0.689	0.034	1464	0.33
1985	0.133	0.028	0.264	0.031	0.471	0.046	0.601	0.043	0.728	0.037	1340	0.34
1986	0.110	0.029	0.226	0.033	0.354	0.046	0.504	0.043	0.736	0.038	1303	0.34
1987	0.123	0.027	0.225	0.031	0.407	0.045	0.597	0.042	0.792	0.036	1296	0.39
1988	0.119	0.028	0.222	0.032	0.378	0.044	0.591	0.045	0.729	0.036	1287	0.36
1989	0.128	0.028	0.211	0.032	0.380	0.043	0.602	0.044	0.757	0.036	1221	0.38
1990	0.131	0.029	0.218	0.034	0.373	0.045	0.587	0.046	0.713	0.037	1189	0.35
1991	0.133	0.030	0.219	0.035	0.388	0.047	0.587	0.045	0.702	0.037	1138	0.37
1992	0.159	0.029	0.254	0.033	0.407	0.043	0.587	0.044	0.755	0.037	1118	0.39
1993	0.195	0.031	0.278	0.035	0.451	0.047	0.613	0.046	0.792	0.038	1158	0.37
1994	0.145	0.031	0.234	0.034	0.336	0.044	0.570	0.043	0.709	0.038	1121	0.37
1995	0.157	0.332	0.263	0.371	0.363	0.046	0.567	0.045	0.684	0.040	1138	0.32
1996	0.184	0.034	0.281	0.038	0.372	0.046	0.660	0.045	0.708	0.041	1120	0.33
1997	0.123	0.036	0.210	0.039	0.281	0.046	0.562	0.046	0.659	0.042	1098	0.32

Source: GSOEP 1984-97, own calculations.

Note: 1) Coefficient of the schooling variable α_1. To obtain the return in percent, compute [exp(α_1)-1] ×100.

Annex 10: Wage premia[1)] associated with education degrees 1984-97 - Women

	Apprentice.		Master		High school		High. tech. coll.		University		Obs.	R^2
	return	s.e.	return	s.e.	return	s.e.	return	s.e.	return	s.e.		
1984	0.196	0.031	0.359	0.054	0.612	0.089	0.567	0.084	0.853	0.067	785	0.27
1985	0.178	0.032	0.321	0.053	0.548	0.089	0.632	0.083	0.860	0.062	758	0.28
1986	0.213	0.033	0.299	0.053	0.527	0.082	0.701	0.099	0.838	0.065	733	0.26
1987	0.197	0.032	0.316	0.050	0.520	0.073	0.591	0.088	0.859	0.060	733	0.29
1988	0.170	0.034	0.280	0.054	0.488	0.074	0.639	0.087	0.860	0.062	712	0.28
1989	0.192	0.036	0.307	0.056	0.470	0.071	0.537	0.087	0.918	0.061	693	0.29
1990	0.189	0.037	0.308	0.056	0.464	0.072	0.563	0.093	0.946	0.063	674	0.29
1991	0.186	0.036	0.329	0.055	0.484	0.072	0.585	0.085	0.942	0.060	715	0.30
1992	0.172	0.033	0.280	0.051	0.404	0.060	0.628	0.078	0.853	0.055	739	0.31
1993	0.191	0.037	0.303	0.057	0.474	0.065	0.549	0.089	0.944	0.624	725	0.30
1994	0.190	0.036	0.277	0.053	0.424	0.064	0.592	0.081	0.888	0.057	736	0.31
1995	0.155	0.039	0.286	0.055	0.462	0.066	0.649	0.081	0.820	0.060	722	0.29
1996	0.173	0.038	0.312	0.052	0.414	0.060	0.607	0.080	0.805	0.059	769	0.25
1997	0.141	0.037	0.264	0.053	0.362	0.058	0.545	0.081	0.794	0.057	746	0.26

Source: GSOEP 1984-97, own calculations.

Note: 1) Coefficient of the schooling variable α_1. To obtain the return in percent, compute $[\exp(\alpha_1)-1] \times 100$.